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Shocks to the international prices of agricultural commodities and the effects on welfare and poverty. A simulation of the ex ante long-run effects for Uruguay☆

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ABSTRACT

In countries with a large share of low- and medium–low-income households, an increase in agricultural commodity prices may have a damaging and widespread effect on the population through the rise in the cost of the consumption basket. A less obvious channel, changes in labor income, would be more beneficial to middle-income households. The simulations show that ex ante all households would experience losses of up to 7.5% of their initial expenditure, with poorer households being the most affected. Poverty would increase by 25% (7.2 p.p.), while indigence would increase even more by 35% (2.25 p.p.). The results also show that, on average, both poor and indigent households would move further away from the threshold lines, meaning that, within each category, poor and indigent households would become more homogeneous. Finally, since Uruguay has a comparative advantage in agricultural commodities, the improvement in the terms of trade is likely to increase the tax revenues, which could be used to compensate those who are negatively affected.

1. Introduction and motivation

Since the beginning of the past decade, there has been an important rise in agricultural commodity prices in international markets; this is particularly noteworthy when considering the first half of 2008 and to a lesser extent the second half of 2011. This phenomenon attracted considerable attention because of the potentially negative impacts on poverty and income distribution through an increase in the prices of staple goods.[1] Most studies have focused on developing and less developed countries, albeit using different dimensions and measures of well-being: welfare, employment, inequality and poverty. This interest was partly driven by the aim of providing policy recommendations to minimize undesirable effects. As a case in point, UNCTAD (2013) devoted one chapter of its Commodities and Development Report to studying the direct effects of the 2003–2011 commodity boom on poverty and food insecurity.

In this paper, we contribute to this literature through the simulation of the ex ante impacts on welfare and poverty arising from an

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[1] Some of the works in this literature are Chávez Martín del Campo et al. (2008), Ivanic and Martin (2008), Porto (2008), Valero-Gil and Valero (2008), De Hoyos and Medvedev (2011), Ferreira et al. (2013), Nelson et al. (2011), Tlhalefang and Galebotswe (2013), Lederman and Porto (2015) and Moncarz et al. (2017).

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increase in the price of agricultural commodities for Uruguay. Two characteristics of Uruguay make it an appealing and relevant case. On the one hand, it is an economy with a large share of households with low and middle–low income, which they spend largely on staple goods. On the other hand, the country enjoys a strong comparative advantage in the production of agricultural goods; this means that, at least at the macro level, there is scope for large benefits from an increase in agricultural commodity prices. Following the standard approach in the literature nowadays, our analysis adopts a micro perspective.

The rest of the paper is organized as follows. Section 2 presents a stylized description of the Uruguayan economy, especially in terms of the variables in which we are most interested. In Section 3, we briefly review the literature related to the nexus between trade policy and poverty.[2] Section 4 introduces a simple theoretical model with the main purpose of underpinning our empirical approach. In Section 5, the empirical framework is laid out. Section 6 discusses the approach adopted to obtain the impact on different measures of well-being through the use of micro-simulations at the household level. Section 7 summarizes the main results.

2. Some stylized figures regarding the Uruguayan economy

During the first decade of the twenty-first century, international commodity prices rose substantially. Agricultural commodities were no exception. In the case of Uruguay, the average agricultural commodity prices increased between 32% and 47% over the period 2002–2011 compared with the 10 preceding years.[3] Fig. 1 shows this situation along with the significant price volatility.

In this context, goods that are intensive in the use of agricultural commodities as inputs, such as staple goods, are likely to experience rising prices. These may be particularly damaging to the low- and medium–low-income groups and ultimately to the population at large if most households fall into these income groups. Table 1 reports the proportion of households with per capita income below different

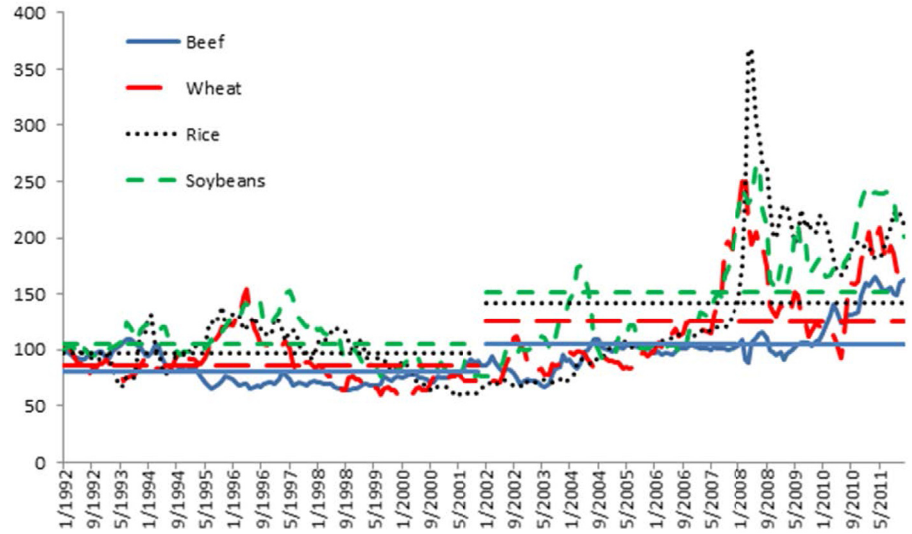

Fig. 1. Prices of the main products exported by Uruguay. 1992–2011. Monthly data.

Table 1
Proportion of households with per capita income lower than value of reference.

	Less or equal than:		
	1/4 of average income	1/2 of average income	Average income
2002	11.0	33.8	67.3
2005	10.0	31.7	67.1
2008	7.9	31.8	67.8
2011	5.9	28.2	66.5

Source: own calculations based on Encuesta Continua de Hogares (Uruguay). In all cases, household income excludes the implicit rent for those households who own the house they live in.

[2] A large body of literature, to which we do not refer here, has focused almost exclusively on less developed countries, where food security is a very important issue, especially for the poorest households.

[3] These figures result from considering agricultural commodities, which explain most of the country's exports.

Table 2
Shares in Consumption and income $^{(\#)}$.

Quintile (*)	FB	CLO	EQU	OTH	Labor income	Salaried labor income
1	39.3	5.1	4.3	51.3	56.4	38.5
2	32.5	3.9	3.6	60.1	62.0	46.3
3	26.6	4.0	3.7	65.7	61.3	47.2
4	23.6	3.8	3.9	68.7	61.8	45.8
5	16.2	3.8	5.3	74.7	57.4	40.2

Source: own calculations based on Encuesta Nacional de Gastos e Ingresos de los Hogares 2005–2006. FB: food and beverages; CLO: clothing; EQU: equipment; OTH: other goods and services. (#) Shares in consumption as a percentage of total household expenditure. (*) Per capita household expenditure.

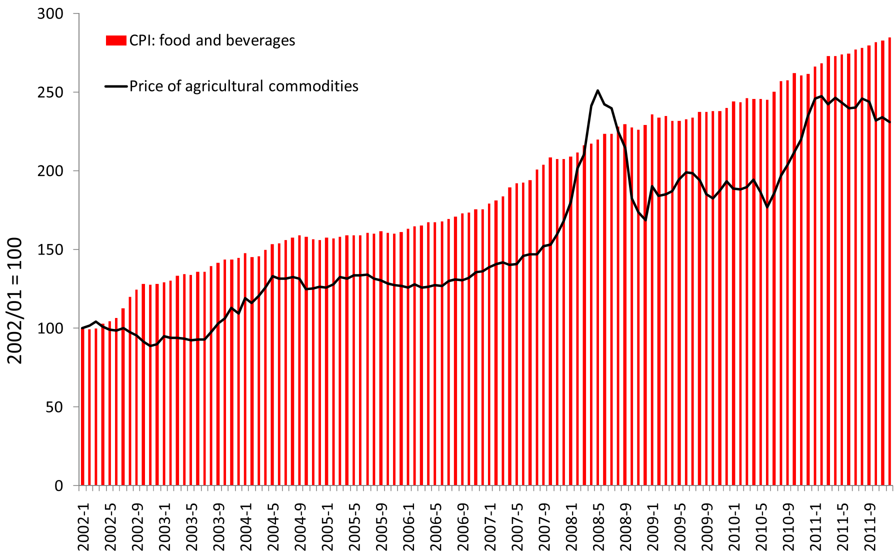

Fig. 2. Agricultural commodity prices and consumer prices of food and beverages: 2000–2011.

thresholds. Firstly, we see a high and persistent number of households with incomes below the average. Around two-thirds of households earn less than the average income; around one-third of households earn less than half the average income; finally, around 10% of Uruguayan households earn less than a quarter of the average income. On the other hand, the latter group – that is, those with extremely low incomes – has diminished in recent years. It is important to note that the values for the early years were heavily influenced by the turmoil that Uruguay experienced due to the collapse of the Argentinean economy in 2002.[4] Another important reason behind the reduction in the share of those with extremely low incomes is the role played by cash transfers from the public sector. Note that, if social transfers are excluded from the calculations, the reduction is significantly smaller, falling from 11.2% in 2002 to 8.3% in 2011 (5.9% in 2011 when including social transfers).

Due to this pattern of income distribution, it is no surprise that a large share of households spend an important proportion of their income on food and beverages (see Table 2), the prices of which, as shown in Fig. 2, moved alongside those of agricultural commodities.

The two patterns just described (a large share of households with a low income and a large proportion of expenditure on staple goods) suggest that there may be potentially large redistributive effects from increases in the prices of agricultural commodities. On the other hand, there may be significant positive macroeconomic effects associated with the increase in the country's terms of trade, which may allow the government some space for redistribution. In the case of Uruguay, redistribution took the form of pro-poor policies: there was a very large increase in the share of social transfers as a fraction of household income, particularly among those households at the bottom end of the income distribution (see Table 3). This increase in transfers took place within a broader set of public policies aimed at having a greater impact on poorer households, such as the provision of health and education services by the public sector. An example of the latter is the increase in health care spending from 3.3% of the GDP in 2005 to 5.4% in 2012.[5]

To sum up, we argue that the case of the Uruguayan economy is relevant to study for two main reasons. On one hand, Uruguay is an economy that is highly exposed to increases in the international prices of agricultural commodities due to its structural patterns of income distribution and expenditure. On the other hand, considering the country's comparative advantages, there is room for important

[4] Even though Uruguay has been successful since then in reducing its dependence on Argentina, and to some extent on Brazil, these economies are still among its main trading partners, especially for the manufacturing sectors.

[5] We thank an anonymous referee for pointing out this issue.

Table 3
Shares of Social Transfers on Household incomes.

	Quintile (household per capita income)				
	1	2	3	4	5
2002	1.9	0.5	0.3	0.1	0.1
2005	4.1	0.8	0.2	0.1	0.1
2008	11.3	3.6	2.2	1.6	1.0
2011	13.6	4.0	2.0	1.2	0.8

Source: own calculations based on Encuesta Continua de Hogares (Uruguay). In all cases, household income excludes the implicit rent for those households who own the house they live in.

benefits from a macro perspective, which could allow the Government the use of compensatory measures directed towards assisting those households that are most negatively affected. These two characteristics may create some interactions that are relevant and insightful to explore. We move onto these issues in the next sections.

3. Related literature

The economic literature on the links between open trade policies and their assumed positive impact on economic growth and development has reached a consensus regarding results that are measured on average. However, because of the broad set of interrelated factors affecting social welfare outcomes because of trade liberalization, when dealing with the potentially beneficial impacts at the household level, this consensus is less obvious. In fact, trade policies have strong redistributive impacts and, in most cases, it is possible to identify winners and losers. Given the particular importance of local institutional arrangements and the way in which markets determine (consumer) local prices, if poor individuals are among those who lose, the long-run opportunities for the development of a country may be compromised.

McCulloch et al. (2001) and Winters et al. (2004) have both contributed to underpinning the scope of the debate. They noted that the empirical evidence, both in cross-country and in country-case-level studies, has so far not provided homogeneous results, with liberalization episodes in which the living conditions of the poorer households declined. A common feature in terms of the choice of the methodology used to assess the direct impact of trade liberalization on poverty is the preference for partial equilibrium techniques instead of general equilibrium (GE) approaches. Indeed, it is crucial to the choice that this approach allows the possibility of identifying household income and consumption effects. A similar analysis applying GE techniques to quantify distributive effects as a result of price shocks will be limited due to the lack of sufficient disaggregation to trace fully the impact of policies on poverty. Of course, the partial equilibrium approach has the drawback that it omits some second-order effects and even some direct effects. Our analysis in this paper shares this drawback.

The literature dealing with the trade liberalization–poverty nexus using the partial equilibrium approach starts with the canonical work of Deaton (1989), which was later expanded and improved with the important methodological contribution by Porto (2006). This approach has been encouraged by the ever-increasing availability of household surveys, especially for developing and less developed countries. For the case of Uruguay, a recent study by Borraz et al. (2012) analyzed the impact of MERCOSUR on poverty and inequality through two main transmission channels: prices and income. They found that, while trade liberalization favored a reduction in poverty indicators, it had almost no effect on income inequality, concluding that trade integration policies cannot be regarded as ‘poverty alleviating’ per se.

Our main goal in this paper is to provide an explanation of how shocks in the prices of agricultural commodities can affect welfare and poverty at the household level. We achieve this by using a country case, namely Uruguay, for the period 2002–2011. We assume an urban model, in which households consume products and sell labor but do not produce agricultural commodities. The latter assumption is not particularly restrictive for the Uruguayan case, since about 91–93% of the population adjusts to our modeling of households. With a few exceptions (i.e. Moncarz et al., 2017 for Argentina), the previous evidence has not dealt directly with the poverty impacts on poorer households of increasing agricultural commodity prices. De Hoyos and Medvedev (2011) analyzed the poverty impact of higher food prices from a global perspective, while Lederman and Porto (2015) offered a recent review of the effects of changes in commodity and other prices on household welfare, with a focus on less developed and developing countries. Finally, we also provide an additional contribution for the case of Uruguay by analyzing the possible redistributive effects associated with an increase in tax collection arising from an improvement in the economy's terms of trade resulting from higher prices of agricultural commodities.

4. Theoretical framework[6]

The theoretical underpinning of our analysis assumes a small open economy that produces and trades S primary commodities, of which $S_A \subset S$ are agricultural commodities. Assuming that the number of primary commodities is at least as large as the number of factors, the factor rewards are fully determined by the commodity prices.

[6] See Moncarz et al. (2017) for a full presentation of the theoretical framework.

$$W = p\left(P_S^D\right) \tag{1}$$

where W is the vector of factor rewards and P_S^D is the vector of commodity prices in the local currency. Since our economy is small, we have:

$$P_S^D = EP_S^*(1+T) \tag{2}$$

where E is the nominal exchange rate, P_S^* is the vector of international commodity prices and T is the vector that reflects the ad valorem equivalent of the country trade policy, so we obtain:

$$W = p\left(P_S^*, E, T\right) \tag{3}$$

There are also M traded manufacturing sectors, of which $M_F \subset M$ produce food goods. Following the current standard approach in the trade literature, the M manufacturing sectors are monopolistically competitive, with production exhibiting increasing returns to scale (IRS). In each m sector, each producer, domestic or foreign, produces a differentiated variety using all the factors of production and primary commodities. In addition, there are N non-traded sectors that are also monopolistically competitive, with each domestic producer producing a differentiated variety under IRS using only the production factors.[7] Given the previous assumptions, the price indices for the traded and non-traded sectors can be expressed as:

$$P_m = \left[N_m\left(p_{i,m}^c\right)^{1-\sigma_m} + N_{m^*}\left(p_{i,m^*}^c\right)^{1-\sigma_m}\right]^{\frac{1}{1-\sigma_m}} \quad \text{for all} \quad m \subset M \tag{4}$$

$$P_n = \left[N_n\left(p_{i,n}^c\right)^{1-\sigma_n}\right]^{\frac{1}{1-\sigma_n}} \quad \text{for all} \quad n \subset N \tag{5}$$

where N_m and N_{m^*} are the number of varieties produced locally and imported respectively by sector m and N_n is the number of non-traded varieties by sector n. $p_{i,j}^c$, for ($j = m$, m^*, n), are the prices in local currency for variety i paid by consumers and produced in j that, under our assumptions, are a function of the factor prices (W and W^*), and σ_j are the elasticity of substitutions.

Using (3), (4) and (5), it can be shown that the prices of manufactured goods are fully determined by the prices of international commodities. These relationships, as well as the effects of international commodity prices on factor prices, are the ones that we aim to estimate. Finally, given the country tax structure, it is possible to derive a relationship between commodity prices and tax revenues.

5. Empirical framework: the pass-through of international prices

5.1. Methodology

An important contribution from a methodological perspective is the work of Porto (2008) on rural Mexico, in which the author extended the methodology originally proposed by Deaton (1990), also identifying the effects taking place through changes in the labor supply of households directly involved in agricultural production. However, even in those cases in which an effort was made to innovate and refine the estimations of the effects on welfare, it is not possible to observe a similar attempt when calculating the pass-through from international prices to domestic ones, in most cases relying on very simple ad hoc rules.

In the present case, two issues need to be faced. First, commodities of which the prices experienced an increase are mostly not directly consumed by households, but their impacts originate because they are used as inputs in the production of staple goods. This is mainly the case of wheat, which is a key ingredient in the production of food goods, especially bakery products and the like. In other cases, like soybeans, these commodities have an important weight in even earlier stages of production, such as the feeding of animals for human consumption. Thus, it is not advisable or appropriate to use simple ad hoc rules that could instead be used in cases in which the goods consumed by households are tradable.

To deal with this issue, there are several alternatives at hand. Firstly, one could perform a cost analysis at the micro level. This would certainly provide a very precise and accurate effect but would be too demanding in terms of data; also, it would require a case (sector) by case (sector) analysis, which would introduce additional problems for obtaining an overall effect on the general level of prices (and even for subcategories, like food and beverages). Related to this, most of the existing research has only considered the effects working through just a few goods, ignoring those that may take place through changes in the prices of other goods that are not directly related to agricultural commodities but certainly may be affected by general equilibrium effects. The same argument applies to the effects on households' income, which, in the vast majority of the literature, has been restricted to that of households that are directly involved in the production of commodities with changing prices. This becomes important in the case of urban households, for which it is straightforward to assume that they do not produce or are not directly involved in the production of agricultural commodities. In light of the above arguments, we propose a more pragmatic and agnostic methodology, obtaining the pass-through of international commodity prices to domestic ones (of goods and factors) using a time series analysis.

[7] The assumption that non-traded goods are produced using only production factors is a simplification that has no impact on the underlying relationships between the international prices of commodities and either consumer prices or factor rewards.

In time series analysis, most of the existing literature on the subject has relied on performing an impulse response analysis to compute the pass-through of international prices to domestic ones. For example, Furlong and Ingenito (1996), Krichene (2008), Zoli (2009), Ferrucci et al. (2010), Rigobon (2010) and Ianchovichina et al. (2014), among others, fitted a Vector Autoregressive (VAR) model and then estimated the corresponding response of domestic prices to a given shock to international commodity prices. However, this approach fails to provide a 'standard' measure of elasticity.

In our case, instead, we propose to estimate the long-run elasticities by identifying a Vector Error Correction (VEC) model in which the international prices of agricultural commodities and the consumer prices are the endogenous variables.[8] An advantage of the VEC approach is that, once an economic structure is introduced into the analysis, it is able to provide the long-run effects of the former over the latter, allowing for a direct estimate of the elasticities according to the usual definition. This approach, which is used in several applications of monetary theory to estimate the pass-through of changes in the exchange rate to inflation, has been almost completely absent from the literature about the welfare effects of the recent increase in agricultural commodity prices.

Let us consider the VEC representation of a VAR of order l, given by:

$$\Delta p_t = \prod p_{t-1} + \Gamma_1 \Delta p_{t-1} + \ldots + \Gamma_{l-1} \Delta p_{t-l+1} \tag{6}$$

Since p_t is a $K \times 1$ vector that contains at least one I(1) variable, Π is a singular $K \times K$ matrix with a rank equal to r. Further, Π can be written as $\Pi = \alpha\beta'$, where the $K \times r$ matrix β is the cointegrating matrix. We are interested in analyzing the $r \times 1$ vector $ec_{t-1} = \beta' p_{t-1}$ containing the cointegration relations between prices. In particular, if the variables included in p_t are expressed in logarithms, the coefficients in β represent the elasticities that measure the response of the domestic prices to the international ones. Provided that the cointegrating rank is known, the reduced-rank maximum likelihood estimator (α_e,β_e) is available; however, it only estimates the cointegrating space consistently. Therefore, it is necessary to identify K-r variables utilizing prior information. We assume that the first part of β is an identity matrix, so it takes the form $\beta' = [Ir{:}\ \beta' k\text{-}r]$, where Ir is an identity matrix of order r, while $\beta' k\text{-}r$ is an rx(K-r) matrix with the coefficients to be identified. For identification purposes, our assumption is that the domestic variables, consumer prices, wages and tax collection, are driven by the prices of international commodities.

As pointed out by Juselius (2006), all the long-run effects are captured in the Π matrix. Given that $\beta' p_{t-1}$ represents a stationary linear combination of the variables, the coefficients of β describe the relations of these variables in the steady state. When these relationships are interpreted according to the economic theory, the concept of cointegration matches the notion of long-run equilibrium. Thus, the methodology allows the computing of the long-run elasticities for our variables of interest with respect to the international prices of agricultural commodities. In particular, we are interested in the following elasticities: consumer prices, factor prices and tax revenues.

5.2. Results

In this section, we present the results for the long-run relationships between: (i) consumer and international commodity prices; (ii) hourly wage rates and international commodity prices; and iii) tax revenues and international commodity prices. For the purposes of the estimation, all the variables are converted into indices, with 2002/01 = 100, and then we take the natural logarithm. The sample period is from January 2002 to December 2011, with monthly data.

With respect to the relationships between the consumer prices and the international prices of agricultural commodities (*pwa*), we group the former into four categories: a) food and beverages (*pcfb*); (b) clothing (*pcclo*); (c) equipment (*pcequ*); and (d) other goods (*pcoth*). Labor, on the other hand, is divided into three mutually exclusive groups – unskilled (*w1*), semi-skilled (*w2*) and skilled (*w3*) – depending on individuals' formal education. Finally, we work with two alternative definitions of tax revenues: total revenues (*rtot*) and value-added tax (*riva*).

Fig. 3 depicts the behavior of the series used to obtain the different elasticities. As a casual inspection shows, an issue to consider is the presence of structural breaks, which could bias the magnitude of the elasticities.[9] As the formal analysis presented below confirms, the commodity prices appear to have experienced a sudden decrease at the beginning of 2009, around the time of the financial crisis that erupted in 2008. Meanwhile, for the domestic series, the breaks seem to have taken place closer to the beginning of the period under study, driven mostly by the severe financial turmoil experienced by Uruguay in the year 2002 as a consequence of the Argentinean economic crisis that started in December 2001 and lasted for most of 2003. A further issue to consider when estimating the different models is the existence of seasonal behavior.

To test the existence of unit roots in the time series, firstly, we follow the procedure suggested by Lanne et al. (2003) to detect a structural change in the data.[10] Secondly, following Lanne et al. (2002), we analyze the existence of unit roots, taking into account the existence of structural shifts in the time series *pcbf*, *pcclo*, *pcequ*, *pcoth* and *pwa* (in the periods in which breaks have previously been identified), by considering three possibilities: a simple shift dummy, a gradual shift and a general nonlinear shift. The results are shown in Appendix 2 (Tables A.1–A.5). For more details on how to run the unit root tests in the presence of structural shifts, see the explanatory

[8] Anderson and Tyers (1992) provided an example of the use of an error correction model to compute elasticities for changes in border prices relative to domestic producer prices. Moncarz et al. (2017), for Argentina, dealt specifically with the impact of the changes in international prices of agricultural commodities on domestic prices.

[9] We thank an anonymous referee for bringing this point to our attention.

[10] It is important to note that the selected unit root test is not sensitive to a slight misspecification of the break dates (Lütkepohl and Krätzig, 2004). The selection of the break date may be sensitive to the lag length. In Appendix 2, the tests for unit roots are computed for different lag lengths according to different model selection criteria: Akaike (AIC), Hannan–Quinn (HQ), final prediction error (FPE) and Schwarz (SC).

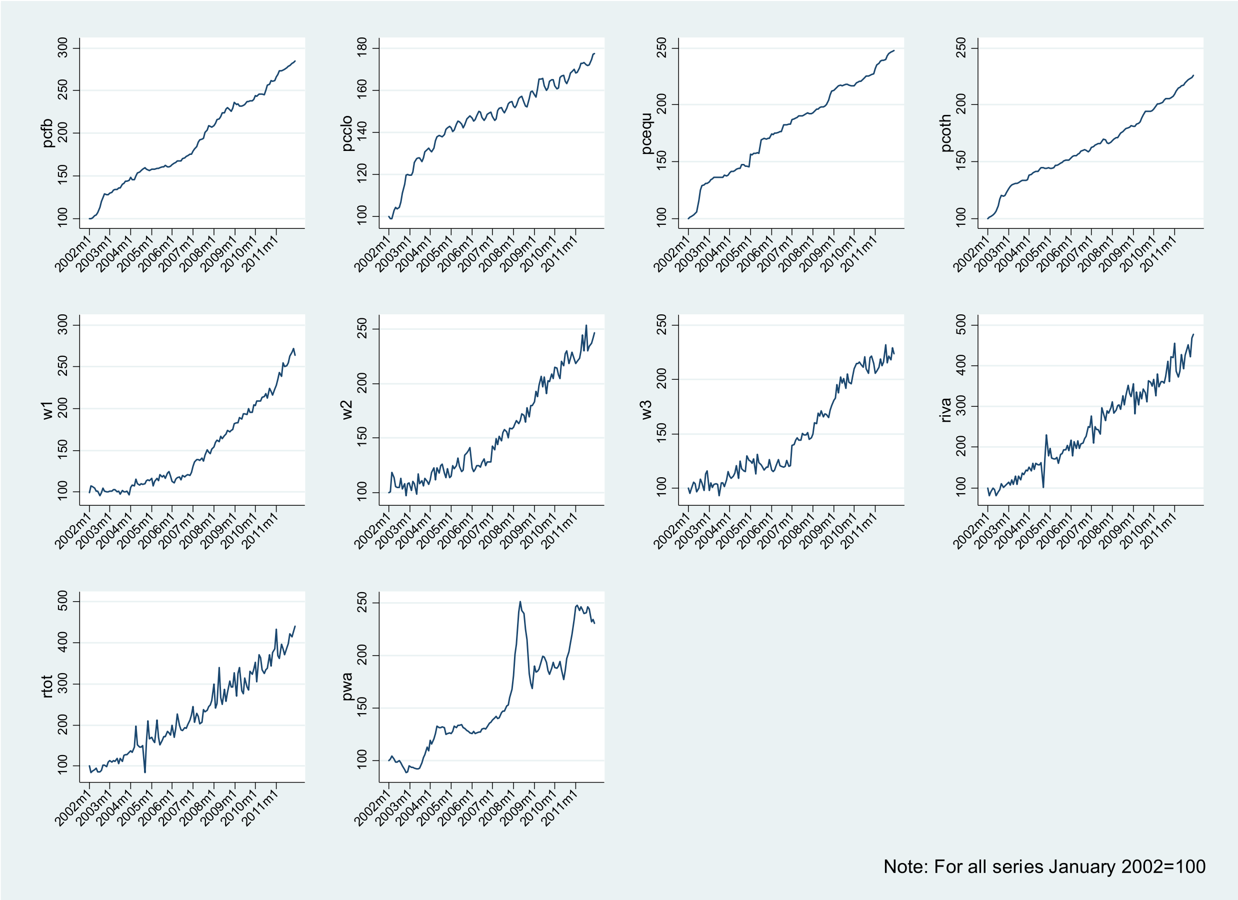

Fig. 3. Time evolution of prices, wages and tax collection.

note in Table A.1. The hypothesis of a unit root cannot be rejected for all the price series. To assess whether there is a (long-run) cointegration relationship between domestic prices and *pwa*, the Saikkonen and Lütkepohl (S&L) (2002) test is performed between *pwa* and each of the consumer prices; dummy variables are included to take into account the presence of the previously identified structural changes.[11] Thus, we run four cointegration analyses between two series, *pwa* and *pcj* (where $j = fb$, *clo*, *equ* and *oth*). The results are presented in Tables A.7–A.10 in Appendix 2 and show that *pwa* is only cointegrated with *pcfb* (the hypothesis that the cointegrating rank is equal to one cannot be rejected), whereas the remaining prices do not seem to be cointegrated with *pwa* (thus, given a system with two variables, $K = 2$, the hypothesis of a cointegrating rank equal to zero cannot be rejected). As *pcclo*, *pcequ* and *pcoth* are not related to *pwa* in the long run, these prices are omitted from the analysis that follows.

Finally, specifying the long-run determinants of *pcfb* correctly requires the consideration of the monetary policy led by the Uruguayan Central Bank (BCU). The reports of the monetary policy of the BCU[12] indicate that, since June 2002, the instrument chosen to control for the inflation has been the quantity of money. This fact is backed by our analysis; Table A.6 in Appendix 2 reports the results of several unit root tests in the presence of a structural shift for the monetary supply. As the results show, the null hypothesis of a unit root for the (log of the) monetary supply (*m1*) cannot be rejected when a structural break is established in the period 2002 M7 after applying the method suggested by Lanne et al. (2003). Table A.11 in Appendix 2 shows the results of the cointegration analysis carried out by taking into account the variables *pwa*, *pcfb* and *m1*. The results indicate that the hypothesis of a cointegrating rank equal to 1 cannot be rejected ($r = 1$). Given that there are three variables (i.e. $K = 3$), the result is that there are $K\text{-}r = 2$ exogenous (permanent) shocks that tend to influence the behavior of *pcfb*. The first shock represents the pushing force associated with *pwa*, and the second one is the stochastic trend driven by the monetary policy of the BCU, represented by *m1*. As shown in Table 4, the elasticity of *pcfb* is positive[13] and statistically different from zero at the usual level of significance.

[11] Saikkonen and Lütkepohl (2000) performed the cointegration tests with level shifts and showed that the limiting distributions of the Johansen-type tests for the cointegrating rank remain unaffected.

[12] http://archivo.presidencia.gub.uy/_web/noticias/2005/06/2005063002.htm

[13] Given that all the variables in the cointegration equations are located on the right-hand side (RHS), the signs of the elasticities should be inverted prior to interpretation.

Table 4
Coefficients of cointegration relations: (*pcfb*, *m1*, *pwa*).

Cointegration Equation	$pcfb_{t\text{-}1}$	$m1_{t\text{-}1}$	$pwa_{t\text{-}1}$	$shiftpcfb_{t\text{-}1}$	$shiftpwa_{t\text{-}1}$
$ec_{1,t\text{-}1}$	1.000	−1.313	−0.623	−0.169	−0.207
	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)

(p-Value). The Final Prediction Error criterion indicated that the optimal VAR lag length (in levels) is equal to 4. The cointegration test was run using S&L Test, including CONST, TREND, SEASONALS and shift dummies indicating breaks in [2002 M7] [2009 M1], named shifteM1 shiftpwa respectively. The null hypothesis H0: rank(β) = 1 cannot be rejected, so that the VEC was specified assuming that the cointegration rank is equal to 1. Remaining VEC's specification details are as follows: deterministic variables: shiftpcfb shiftpwa CONST S1 S2 S3 S4 S5 S6 S7 S8 S9 S10 S11 TREND, endogenous lags (in differences): 3, sample range: [2002 M5, 2011 M12], T = 116, estimation procedure: One stage. Johansen approach. Source: Own calculations.

Table 5
Coefficients of cointegration relations: wages.

Cointegration Equation	$w1_{t\text{-}1}$	$w2_{t\text{-}1}$	$w3_{t\text{-}1}$	$m1_{t\text{-}1}$	$pwa_{t\text{-}1}$
$ec_{1,t\text{-}1}$	1.000	0.000	0.000	1.521	−0.222
	(0.000)	(0.000)	(0.000)	(0.000)	(0.039)
$ec_{2,t\text{-}1}$	0.000	1.000	0.000	0.667	−0.144
	(0.000)	(0.000)	(0.000)	(0.002)	(0.031)
$ec_{3,t\text{-}1}$	0.000	0.000	1.000	0.550	−0.172
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

(p - Value). The Schwarz and Hannan-Quinn Criteria indicated that the optimal VAR lag length (in levels) is equal to 1. The cointegration test was run using S&L Test, including CONST, TREND, SEASONALS and shift dummies indicating breaks in [2002 M7] [2009 M1] [2004 M10], named shifteM1, shiftpwa and shiftriva respectively. The null hypothesis H0: rank(β) = 3 cannot be rejected, so that the VEC was specified assuming that the cointegration rank is equal to 3. Remaining VEC's specification details are as follows: deterministic variables: shiftw3 shiftpwa CONST S1 S2 S3 S4 S5 S6 S7 S8 S9 S10 S11 TREND, endogenous lags (in differences): 0, sample range: [2002 M2, 2011 M12], T = 119, All the deterministic terms were included in the cointegrating relationships. Only coefficients for M and *pwa* are showed. Estimation procedure: One stage. S2S approach. Source: Own calculations.

In the case of the hourly wage rates, we work with three categories of labor, unskilled (*w1*), semi-skilled (*w2*) and skilled (*w3*). The tests for unit roots allowing for the presence of structural shifts show that the three variables are I(1) according to the results reported in Tables A.12–A.14 in Appendix 2, and the test for cointegration does not reject the null of three cointegrating relationships, as reported in Tables A.15 and A.16 in Appendix 2 (here also the Lütkepohl and Saikkonen (2000) test is employed, adding dummies to indicate the presence of structural change). As before, we assume that the international prices are exogenous. The results, reported in Table 5, show that, for the three categories of labor, the elasticities with respect to the international prices are positive, with slightly larger values for unskilled labor, and all the estimates are statistically significant at the 5% level.

Finally, we estimate two additional models, one for the tax revenue of the value-added tax (VAT), *riva*, and another for the total tax collection, *rtot*, also including the variables *pwa* and *m1*. As in the previous cases, the tests for unit roots do not reject the null of the series being I(1), as reported in Tables A.17 and A.19 in Appendix 2. Tables A.18 and A.20 in Appendix 2 report that, in both cases, the null of one cointegrating relationship is not rejected. Table 6 shows, as could be expected, that the elasticities with respect to the international prices are positive and statistically significant. The estimate is higher for the case of the VAT, which is not at all surprising, since the total tax collection includes some revenues that are at best weakly related to international prices.

6. Simulation of welfare and poverty following an increase in the international prices of agricultural commodities

6.1. Welfare effects

Our primary goal is to simulate the long-run effects on welfare that may follow a permanent increase of the international prices of agricultural commodities. Once we have obtained the elasticities of consumer prices, wages and tax collection with respect to the international prices of agricultural commodities, we can then simulate the welfare effects that would follow an increase in the latter.

In particular, the welfare effect on household *h* is measured by the negative of the compensating variation relative to its initial expenditure. Taking into account the fact that only the prices of food and beverages appear to be influenced by changes in the prices of agricultural commodities, we have:

$$\frac{dx_0^h}{e^h} = -s^h \psi_{p_{s_A}} d\ln p_{s_A} + \left(\sum_j \theta_j^h \varepsilon^j_{w,p_{s_A}}\right) d\ln p_{s_A} + TR_m^h + TR_q^h \tag{7}$$

where s^h is the budget share spent on varieties produced by the food and beverage sector, $\psi_{p_{sA}}$ is the elasticity with respect to the international price index of agricultural commodities (p_{sA}) of the consumer price index of food and beverages, θ_j^h is the salaried labor income of member *j* as a share of the total income of household *h* and $\varepsilon^j_{w,p_{sA}}$ is the wage elasticity that captures the proportional change in the wage rate of household member *j* as a response to the change in the international prices of agricultural commodities. Finally, TR_m^h is

Table 6
Coefficients of cointegration relations: tax collection.

	VAT collection (*)	Total tax collection (**)
tax_coll_{t-1}	1.000	1.000
	(0.000)	(0.000)
M_{t-1}	−1.077	−0.530
	(0.001)	(0.000)
pwa_{t-1}	−0.405	−0.160
	(0.000)	(0.057)
$shiftriva_{t-1}$	−0.117	−0.092
	(0.000)	(0.008)
$shiftem1_{t-1}$	−0.179	−0.146
	(0.001)	(0.001)
$shiftpwa_{t-1}$	−0.063	−0.029
	(0.210)	(0.466)
$trend_{t-1}$	0.010	−0.001
	(0.086)	(0.678)

(p - Value). (*) The Akaike criterion indicated that the optimal VAR lag length (in levels) is equal to 4. The cointegration test was run using S&L Test, including CONST, TREND, SEASONALS and shift dummies indicating breaks in [2002 M7] [2009 M1] [2004 M10], named shifteM1, shiftpwa and shiftriva respectively. The null hypothesis H0: rank(β) = 1 cannot be rejected, so the VEC was specified assuming that the cointegration rank is equal to 1. Remaining VEC's specification details are as follows: deterministic variables: shiftriva shifteM1 shiftpwa CONST S1 S2 S3 S4 S5 S6 S7 S8 S9 S10 S11 TREND, endogenous lags (in differences): 3, sample range: [2002 M5, 2011 M12], T = 116, estimation procedure: One stage. Johansen approach. (**) The Akaike, Hannan-Quinn, Schwarz, and Final Prediction Error criteria indicated that the optimal VAR lag length (in levels) is equal to 1. The cointegration test was run using S&L Test, including CONST, TREND and shift dummies indicating breaks in [2002 M7] [2009 M1] [2004 M10], named shifteM1, shiftpwa and shiftriva respectively. The null hypothesis H0: rank(β) = 1 cannot be rejected, so the VEC was specified assuming that the cointegration rank is equal to 1. Remaining VEC's specification details are as follows: deterministic variables: shiftriva shifteM1 shiftpwa CONST TREND, endogenous lags (in differences): 1, sample range: [2002 M2, 2011 M12], T = 119, estimation procedure: One stage. S2S approach. Source: Own calculations.

the cash transfer received by household h as a proportion of its initial expenditure, while TR_q^h represents the participation of household h in the increase in education and health public expenditures. Both TR_m^h and TR_q^h are assumed to be financed with the increase in tax collection due to the improvement in the macroeconomic performance that can be expected for a land-abundant economy such as Uruguay. The assumption behind the use of TR_q^h is that the rise in the expenditures on education and health means an increase in the quantity and/or quality of such services, so they can be considered as in-kind transfers; otherwise, they would have no welfare impact on households.

The first term on the RHS of (7) is the welfare effect that takes place through consumption, while the second term measures the effect through changes in labor income. Considering the way in which equation (7) is computed, a negative value means a welfare loss, while a positive value means a welfare gain. In equation (7) we do not consider second-order effects that take place through changes in consumption patterns in response to changes in consumer prices or changes in the supply of labor. In addition, because of the lack of appropriate data, we do not take into account the effects on non-labor income or those due to the consumption of own-produced goods. Finally, as already mentioned above, all the following analyses ignore any impact on the rural population, which is a relatively minor percentage in the case of Uruguay.

In the simulations of the welfare effects, we need to deal with two sources of randomness. The first comes from the sampling variability of expenditure shares and of the participation of salaried labor in household income, while the second emerges because of the error associated with the estimation of the different elasticities with respect to the agricultural commodity prices. The first source of randomness is controlled by weighting each observation by the inverse probability of its inclusion in the sample, which is provided by the household survey with which we work. For the second source, we follow Porto (2006) and resample 200 times from the empirical asymptotic distribution of the estimated coefficients obtained from the VEC models. After the 200 replications, we compute the standard errors to build the confidence bands.[14]

Before commenting on the results, it is important to stress once again that we are not providing an estimate of the ex post changes in our measures of interest following the rise in commodity prices during the period 2002–2011 but simulating the ex ante long-run impacts of such an increase. For the design and implementation of the necessary compensatory measures, if these were indeed needed, having a proper approximation of the magnitude of the expected effects is important.

Fig. 4 reports the long-run effects on welfare after a permanent increase of 50% in the prices of the main agricultural commodities exported by Uruguay. Unsurprisingly, because of the greater weight of food and beverages in the consumption basket of poorer households, these are the most negatively affected. This pattern is clearly visible in the left panel of Fig. 4, in which the welfare losses through the consumption side vary between 3% and 11% of the initial expenditure, with households at the lower end of the expenditure per capita distribution being the most affected.

Regarding the positive effects following the increase in salaried labor income, the welfare gains vary between 2.5% and 4% of the initial expenditure with an inverted u-shaped distribution (see the middle panel of Fig. 4). In this case, middle-class households benefit

[14] The dashed lines are the 90% confidence intervals.

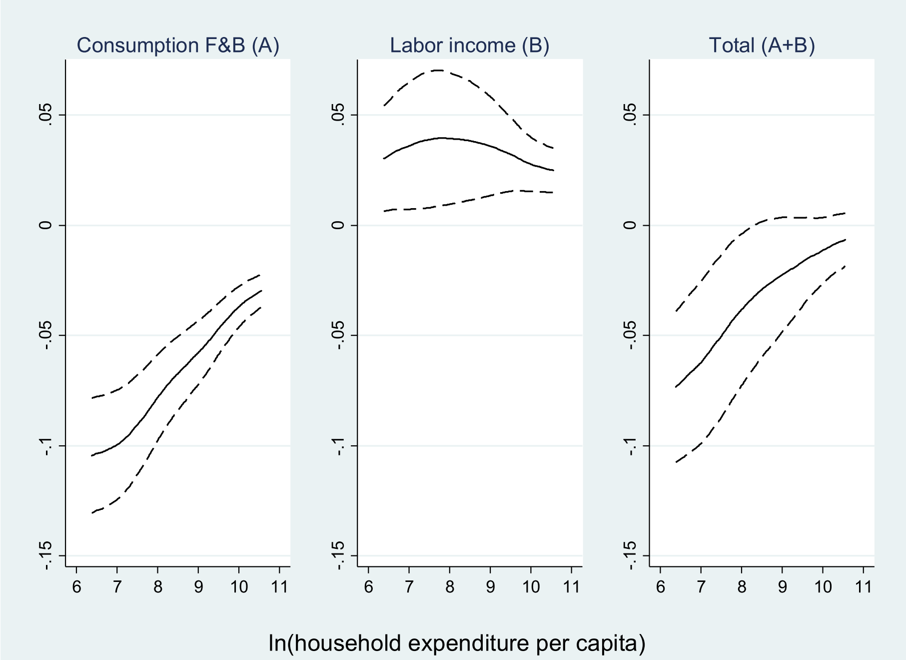

Fig. 4. Welfare effects: Consumption and Labor income (#).

slightly more than households with lower incomes do. Another result is that there are small differences among households. Two explanations for this finding can be distinguished. Firstly, the similar values obtained for the wage elasticities for the three different types of labor tend to reduce the income differences across families; secondly, unlike the case of the consumption of food and beverages, the contribution to the total income made by the salaried labor is less heterogeneous when we compare poor and rich households.

Once we add the consumption and income effects, the first prevails over the second, such that, as reported in the right panel of Fig. 4, all households experience a welfare loss, which decreases with the level of expenditure per capita. The losses are just about 1% in the case of the richest households, reaching around 7.5% of the initial expenditure of the poorest ones. However, for households with a level of expenditure per capita in the upper half of the distribution, we cannot rule out the possibility that the effects are not significant.

As mentioned before, an increase in agricultural commodity prices has the potential to favor the Uruguayan economy overall, given its patterns of comparative advantages in such commodities. These positive impacts at the macro level are expected to be translated into an increase in the country's GDP and consequently in the Government's capacity to increase its tax revenues. As shown in the previous section, this was indeed the case during the period analyzed here. To account for this issue, we simulated the increase in total tax revenues following our assumed 50% increase in the international prices of agricultural commodities. Using the elasticity estimated for the total tax collection and the year 2006 as the reference, the increase in tax revenues in the long run would be about USD 290 million, which is an 8% increase.

To calculate the amount and distribution of new transfers, we work as follows. Under the assumption that the Government uses the extra tax collection in the same way as before the increase in commodity prices,[15] we obtain the increase that would take place in different types of social expenditure: total social expenditure, social assistance (mostly directed to poorer households), retirement and pension benefits and expenditures on education and health by the public sector. For the cases of total social expenditure, retirement and pension benefits and expenditures on education and health, the extra expenditure is distributed among households using the distribution among the deciles of household per capita income reported by Llambí et al. (2010); within each decile, each household receives an identical amount. We assume that only households belonging to percentiles 1 to 40 of the per capita household income distribution receive social assistance transfers, and we assign an identical transfer to each household.

Fig. 5 shows the welfare effects after the distribution of the four types of transfers compared with the situation without such transfers. On the one hand, the households located at the low end of the per capita expenditure distribution benefit the most from the disbursements in social assistance, education and health. On the other hand, the opposite result arises for retirement and pension benefits. The latter outcome is not surprising, since it is more likely that members of middle- and upper-income households are able to meet the conditions required for a retirement benefit.

Taken one by one, with the exception of retirement and pension benefits for households at the upper end of the household expenditure per capita distribution, none of the four kinds of transfers alone can compensate for the negative effects derived from the

[15] This assumption means that our simulations should be understood as the minimum expected effect considering that the Government carried out a policy aimed at increasing the importance of social expenditure in the total expenditure.

Fig. 5. Welfare effects: before and after transfers and education and health expenditures (#).

increase in commodity prices; in all cases, the dotted lines representing the after-transfer situation show negative values. However, when we consider social expenditure as a whole (see Fig. 6), households above the mean of the household expenditure per capita are better off than before the increase in commodity prices; for the remaining households, the losses reduce considerably.

Finally, an important element to take into account regarding our simulations is the fact that the amount of transfers financed with the increase in tax revenues is still below the change in the latter. For instance, taking once again the figures for 2006 as a reference, the increase in transfers would amount to about USD 169 million (when considering the entire social expenditure), while the increase in tax

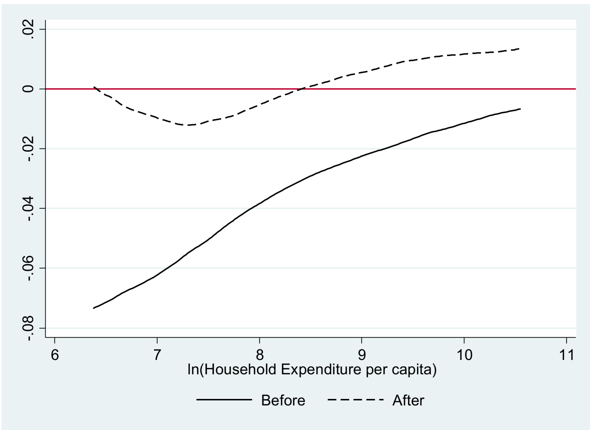

Fig. 6. Welfare effects: before and after social expenditure (#).

Table 7
Indigence and poverty rates before transfers.

Variable			Value	Std. Error (*)
Indigence	Rate	Pre	6.37	0.34
		Post (a)	8.62	0.45
	Gap	Pre	2.07	0.14
		Post (a)	3.04	0.22
	Severity	Pre	1.08	0.10
		Post (a)	1.58	0.13
Poverty	Rate	Pre	28.65	0.69
		Post (a)	35.86	0.81
	Gap	Pre	11.68	0.34
		Post (a)	15.28	0.42
	Severity	Pre	6.53	0.23
		Post (a)	8.76	0.28

(a) After a permanent 50% increase in international prices of agricultural commodities. (*) Bootstrapped standard errors.
Source: author's calculations.

revenues would amount to about USD 290 million. Moreover, the estimated losses for all households are about USD 130 million, which is also less than the USD 290 million increase in tax revenues. These differences leave enough room for a more aggressive compensatory policy if the Government decided to adopt one.

6.2. Indigence and poverty

Although the previous analysis is quite illustrative, because it offers a view of the whole distribution and not just some summary measures, the notion of welfare change as measured by compensatory variation is neither easy nor straightforward to interpret. Thus, to complement the previous results, we also simulate the effects that a 50% increase in the international prices of agricultural commodities has on indigence and poverty by calculating the rate, the gap and the severity of both indicators.[16] A more practical reason for using these two indices is that they are widely referred to when analyzing the living conditions of households, especially in developing countries, where absolute instead of relative poverty is still a persistent phenomenon.

In Uruguay, indigence and poverty are measured in absolute terms, comparing households' income with the minimum expenditures required to avoid falling into either of the two categories. More specifically, the indigence line is defined in per capita terms and built on the basis of the food requirements of a group of representative households according to the socioeconomic and demographic characteristics of an individual of reference, measuring the minimum expenditure necessary to acquire the basic food basket (CBA), which is calculated to guarantee the intake of a certain number of calories. On the other hand, the poverty line is obtained by multiplying the CBA by Engel's coefficient, giving the total basic basket (CBT). Finally, multiplying the CBA and the CBT by the household size, the indigence and poverty baskets are calculated for each household.

To obtain the new value of the CBA for each household, we use the elasticity of the price index of food and beverages with respect to the international commodity prices and update the original indigence line for the time when each household was surveyed.[17] Then the new poverty line is obtained using Engel's coefficient for the time when each household was surveyed. With respect to the new household incomes, these are calculated taking into account only the effect on the labor income of salaried household members and potentially the increase in cash transfers received from the Government.

Table 7 reports the indigence and poverty rates in 2006 and those that would follow in the long-run after a permanent increase of 50% in the prices of agricultural commodities as well as the gap and severity of both variables. The simulations show that indigence increases by 2.25 p.p. and poverty by 7.21 p.p. The relative increase is about 25% for poverty and 35% in the case of indigence.

Another interesting result is that, if we consider the depth of indigence and poverty instead of using a headcount measure, the gap and severity of both measures, especially in the case of indigence, increase more than the corresponding rates. In other words, in addition to an increase in indigence and poverty in response to the rise in the international prices of agricultural commodities, households that were already in those states, as well as those entering them, tend to move further away from the threshold lines. This result means that indigent/poor households tend to become more homogeneous groups, while they become more heterogeneous relative to those with an income above the threshold lines.

Once we account for the cash transfers (social assistance and retirement benefits) financed with part of the increase in tax revenues, and considering the outcomes of the welfare analysis, the increases in indigence and poverty are lower than those in the scenario without transfers; still, both variables experience a rise. As reported in Table 8, for the most favorable case, the increase in indigence is 2.01 p.p. and for poverty it is 6.86 p.p. These figures rise to 2.07 p.p. and 7.05 p.p. when we consider only transfers directed to social

[16] The rate, gap and severity are calculated following Foster et al. (1984).
[17] It would have been more appropriate to work with the changes in the prices of goods that constitute the CBA, data to which we do not have access; however, the correlation of the consumer price index for food and beverages with the indigence and poverty lines is around 0.99 in both cases.

Table 8
Indigence and poverty rates after transfers.

			Value	Std. Error (*)
Indigence	Rate	Post (b)	8.44	0.45
		Post (c)	8.38	0.41
	Gap	Post (b)	2.94	0.20
		Post (c)	2.90	0.19
	Severity	Post (b)	1.51	0.14
		Post (c)	1.49	0.12
Poverty	Rate	Post (b)	35.70	0.73
		Post (c)	35.51	0.81
	Gap	Post (b)	15.09	0.44
		Post (c)	14.93	0.41
	Severity	Post (b)	8.61	0.29
		Post (c)	8.52	0.30

(b) After a permanent 50% increase in international prices of agricultural commodities, and social assistance transfers only to percentiles 1 to 40.
(c) After a permanent 50% increase in international prices of agricultural commodities, and retirement and social assistance transfers. The latter only to percentiles 1 to 40.
(*) Bootstrapped standard errors.
Source: author's calculations.

assistance. A result that still prevails is that the changes in the gap and severity of indigence and poverty are proportionally larger than the corresponding changes in rates, so households falling into either of the two categories move further away from the threshold lines.

The previous results are in line with those of the welfare analysis. Households at the lower end of the per capita expenditure distribution, which are more likely to fall into indigence and/or poverty, are the most negatively affected even after weighing in the positive effects of in-cash transfers. In addition, the inclusion of retirement benefits has a small role, since in most cases these transfers are directed to households that are neither indigent nor poor.

Finally, we need to recall that our results should be taken with some caution. As pointed out previously, the amount of transfers implied by our simulations is well below the expected increase in tax revenues. Thus, for an economy like Uruguay, with strong comparative advantages in agricultural commodities, an improvement in its terms of trade has the potential to generate enough resources to compensate those who otherwise would be hurt. Finally, once we account for the effects of the public social expenditure as a whole—not only monetary transfers that only have an impact on indigence and poverty but also in-kind transfers and public services such as education and health[18]—the results improve substantially compared with those obtained from only considering monetary compensations.

7. Summary and conclusions

This paper deals with the ex ante long-run effects on welfare, poverty and indigence that would follow a permanent increase in the international prices of agricultural commodities for the case of Uruguay for the period 2002–2011. The welfare effect is measured by the negative of the compensatory variation with respect to the initial expenditure. To perform the simulations, we estimate the long-run elasticity for the price index of food and beverages with respect to agricultural commodity prices (other categories of goods appear not to be cointegrated with international commodity prices). We follow a similar procedure for the wage rates of three categories of salaried labor (skilled, semi-skilled and unskilled) and for the collection of taxes. The elasticities were obtained through the estimation of different VEC models, which allows the cointegration coefficients to be interpreted as elasticities.

The simulations show a relative increase of around 25% in poverty and 35% in the case of indigence. Considering the gap and depth of indigence and poverty, we show that these increase even more than the corresponding rates, especially in the case of indigence. In other words, in addition to an increase in indigence and poverty in response to rising international prices of agricultural commodities, households that were already poor and/or indigent (as well as those entering those states) tend to move further away from the threshold lines. Thus, the results suggest that these groups tend to become more homogeneous within themselves and more heterogeneous with respect to households with incomes above the threshold lines. In other words, the distinction between included and excluded households becomes more severe.

The effects on indigence and poverty are smaller but still show an increase when considering cash transfers financed with increasing tax revenues. In the most favorable scenario, the increase in indigence is 2.01 p.p., and for poverty it is 6.86 p.p.; these values increase to 2.07 p.p. and 7.05 p.p. when considering only transfers directed to social assistance. In either case, transfers are more effective, but just slightly, in reducing extreme poverty. The reason behind this outcome is that most households receiving retirement benefits are neither indigent nor poor.

Because of the lack of appropriate data, incomes earned from land owning and from participation in commercial business cannot be incorporated into the analysis. Our study only considers labor income for salaried workers, but for higher-income households land rents

[18] In kind transfers, especially the provision of public services such as education and health are important when indigence and poverty are measured using a multidimensional approach.

and business profits may be an important component of their income, while for low-income households non-salaried labor may have more weight.

Finally, for an economy like Uruguay, with strong comparative advantages in the production of agricultural commodities, another potentially beneficial impact is the increasing capability of the Government to finance social expenditure, not only through cash transfers but also through in-kind transfers and the provision of public services such as health and education, which tend to favor poorer households.

Appendix 1. Data description and sources

Agricultural commodity prices: weighted average of the country main agricultural exports (Beef, Wheat, Rice and Soybeans). Exports values are used as weights. Price indices are from www.indexmundi.com while exports are from WITS of World Bank.

Consumer Prices: are from Instituto Nacional de Estadística.

Wages: average wages are from Instituto Nacional de Estadística using the Encuesta Continua de Hogares.

Nominal Exchange Rate: is from Banco Central del Uruguay.

Household expenditures and incomes: are from Instituto Nacional de Estadística, Encuesta Nacional de Gastos e Ingresos de los Hogares 2005–2006.

Tax collection: Dirección General Impositiva.

Social expenditure: Ministerio de Desarrollo Social.

Appendix 2. Unit root and Cointegration tests

Table A.1
Unit root tests in the presence of structural shift for *pcfb*

Shift function	Test Statistic	Lags (1st diff)	Break date	Critical values		
				10%	5%	1%
$f_t^{(1)'}\gamma$	−1.6275	0 (AIC, FPE, HQ, SC)	2002 M8	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−1.8432	0 (AIC, FPE, HQ, SC)	2002 M8			
$f_t^{(3)}(\theta)'\gamma$	2.9413	0 (AIC, FPE, HQ, SC)	2002 M8			

Note: Time trend included. Critical values (Lanne et al., 2002). Lanne et al. (2002) that analyze the existence of unit roots in the following model: $y_t = \mu_0 + u_1 t + f_t(\theta)'\gamma + x_t$. Where θ and γ are unknown parameters, while x_t is a *AR(p)* process. This expression introduces a shift function $f_t(\theta)'\gamma$ that can be added to the deterministic term μ_0 (see Lütkepohl and Krätzig (2004) for more details). The shift function could represent a simple shift dummy, so that $f_t(\theta)'\gamma = f_t(\theta)'\gamma = f_t^{(1)'}\gamma$ (in this case there is no need of an extra parameter θ, and the function is a simple shift dummy variable with shift date T_B). The second case refers to one in which $f_t(\theta)'\gamma = f_t^{(2)}(\theta)'\gamma$ is an exponential function that depicts a gradual shift to a new level beginning at time T_B. Finally, the third possibility that the test offers to tests unit roots with structural change is to consider a general nonlinear shift, $f_t(\theta)'\gamma = f_t^{(3)}(\theta)'\gamma$. This test must be executed by knowing *a priori* the shift date T_B and the number of lags.

Table A.2
Unit root tests in the presence of structural shift for *pcclo*

Shift function	Test Statistic	Lags (1st diff)	Break date	Critical values		
				10%	5%	1%
$f_t^{(1)'}\gamma$	−0.8176	0 (SC, HQ)	2002 M8	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−0.9205	0 (SC, HQ)	2002 M8			
$f_t^{(3)}(\theta)'\gamma$	−1.3882	0 (SC, HQ)	2002 M8			
$f_t^{(1)'}\gamma$	−1.4173	1 (AIC, FPE)	2002 M11			
$f_t^{(2)}(\theta)'\gamma$	−1.4312	1 (HQ, SC)	2002 M11			
$f_t^{(3)}(\theta)'\gamma$	−1.0290	1 (AIC, FPE, HQ, SC)	2002 M11			

Note: Time trend and seasonal dummies included. Critical values (Lanne et al., 2002). For more details on how to run the unit root tests in the presence of structural shift see footnote in Table A.1.

Table A.3
Unit root tests in the presence of structural shift for *pcequ*

Shift function	Test Statistic	Lags (1st diff)	Break date	Critical values		
				10%	5%	1%
$f_t^{(1)'}\gamma$	−1.5348	0 (SC, FPE)	2002 M7	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−1.2073	0 (SC, FPE)	2002 M7			
$f_t^{(3)}(\theta)'\gamma$	−1.1755	0 (SC, FPE)	2002 M7			
$f_t^{(1)'}\gamma$	−0.6064	6 (AIC, FPE)	2005 M7			
$f_t^{(2)}(\theta)'\gamma$	−0.6265	6 (AIC, FPE)	2005 M7			
$f_t^{(3)}(\theta)'\gamma$	−0.3217	6 (AIC, FPE)	2005 M7			

Note: Time trend included. Critical values (Lanne et al., 2002). For more details on how to run the unit root tests in the presence of structural shift see footnote in Table A.1.

Table A.4
Unit root tests in the presence of structural shift for *pcoth*

Shift function	Test Statistic	Lags (1st diff)	Break date	Critical values		
				10%	5%	1%
$f_t^{(1)'}\gamma$	−1.0062	2 (SC)	2004 M1	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−0.9991	2 (SC)	2004 M1			
$f_t^{(3)}(\theta)'\gamma$	−0.5275	2 (SC)	2004 M1			
$f_t^{(1)'}\gamma$	−1.2400	6 (AIC, FPE, HQ)	2004 M1			
$f_t^{(2)}(\theta)'\gamma$	−1.2528	6 (AIC, FPE, HQ)	2004 M1			
$f_t^{(3)}(\theta)'\gamma$	−1.5546	6 (AIC, FPE, HQ)	2004 M1			

Note: Time trend included. Critical values (Lanne et al., 2002). For more details on how to run the unit root tests in the presence of structural shift see footnote in Table A.1.

Table A.5
Unit root tests in the presence of structural shift for *pwa*

Shift function	Test Statistic	Lags (1st diff)	Break date	Critical values		
				10%	5%	1%
$f_t^{(1)'}\gamma$	−3.3603	2 (HQ, SC)	2009 M1	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−3.3320	2 (HQ, SC)	2009 M1			
$f_t^{(3)}(\theta)'\gamma$	−3.4400	2 (HQ, SC)	2009 M1			
$f_t^{(1)'}\gamma$	−3.7378	4 (AIC, FPE)*	2009 M1			
$f_t^{(2)}(\theta)'\gamma$	−3.6855	4 (AIC, FPE)*	2009 M1			
$f_t^{(3)}(\theta)'\gamma$	−3.8477	4 (AIC, FPE)	2009 M1			
$f_t^{(1)'}\gamma$	−2.6561	6 (AIC, FPE)	2009 M1			
$f_t^{(2)}(\theta)'\gamma$	−2.6564	6 (AIC, FPE)	2009 M1			

Note: Time trend included. Critical values (Lanne et al., 2002). *Results when performing the criteria for an unrestricted VAR. For more details on how to run the unit root tests in the presence of structural shift see footnote in Table A.1.

Table A.6
Unit root tests in the presence of structural shift for *m1*

Shift function	Test Statistic	Lags (1st diff)	Break date	Critical values		
				10%	5%	1%
$f_t^{(1)'}\gamma$	−1.9466	2 (AIC, FPE, HQ)	2010 M3	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−1.9014	2 (AIC, FPE, HQ)	2010 M3			
$f_t^{(3)}(\theta)'\gamma$	−2.5909	2 (AIC, FPE, HQ)	2010 M3			
$f_t^{(1)'}\gamma$	−2.6690	0 (SC)	2002 M7			
$f_t^{(2)}(\theta)'\gamma$	−2.6443	0 (SC)	2002 M7			
$f_t^{(3)}(\theta)'\gamma$	−2.8207	0 (SC)	2002 M7			

Note: Time trend and seasonal dummies included. Critical values (Lanne et al., 2002). For more details on how to run the unit root tests in the presence of structural shift see footnote in Table A.1.

Table A.7
Test for cointegrating rank for (*pwa*, *pcfb*).

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L	3 (AIC, FPE, HQ, SC)	r = 0	14.80	0.0718	13.88	15.76	19.71
		r = 1	2.07	0.5218	5.47	6.79	9.73
S&L with shift dummy[a]	3 (AIC, FPE, HQ, SC)	r = 0	14.13	0.0916	13.88	15.76	19.71
		r = 1	1.67	0.6168	5.47	6.79	9.73
S&L with shift dummy[b]	3 (AIC, FPE, HQ, SC)	r = 0	13.46	0.0101	8.18	9.84	13.48

Note: Deterministic terms in all models: constant, linear trend. Sample period: 2002 M1–2011 M12 (including pre sample values). Critical Values from Lütkepohl and Saikkonen (2000).
[a] Shift(s): [2009 M1] [2002 M8].
[b] Trend orthogonal to cointegration relation.

Table A.8
Test for cointegrating rank for (*pwa*, *pcclo*).

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L	3 (AIC, FPE, HQ)	r = 0	13.27	0.1236	13.88	15.76	19.7
		r = 1	0.36	0.9505	5.47	6.79	9.73
	2 (SC)	r = 0	9.09	0.4268	13.88	15.76	19.71
		r = 1	0.33	0.9577	5.47	6.79	9.73

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Table A.8 (*continued*)

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L with shift dummy[a]	3 (AIC, FPE, HQ)	r = 0	12.28	0.1717	13.88	15.76	19.71
		r = 1	0.81	0.8425	5.47	6.79	9.7
	2 (SC)	r = 0	9.28	0.4076	13.88	15.76	19.71
		r = 1	0.27	0.9688	5.47	6.79	9.73
S&L with shift dummy[b]	3 (AIC, FPE, HQ)	r = 0	8.26	0.0970	8.18	9.84	13.48
	2 (SC)	r = 0	8.38	0.0924	8.18	9.84	13.48

Note: Deterministic terms in all models: constant, linear trend and seasonal dummies. Sample period: 2002 M1–2011 M12 (including pre sample values). Critical Values from Lütkepohl and Saikkonen (2000).

[a] Shift(s): [2009 M1] [2002 M8].

[b] Trend orthogonal to cointegration relation.

Table A.9

Test for cointegrating rank for (*pwa*, *pccequ*).

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L	3 (AIC, FPE)	r = 0	12.67	0.1513	13.88	15.76	19.71
		r = 1	0.10	0.9931	5.47	6.79	9.73
	2 (HQ, SC)	r = 0	8.04	0.5428	13.88	15.76	19.71
		r = 1	0.93	0.8098	5.47	6.79	9.73
S&L with shift dummy[a]	3 (AIC, FPE)	r = 0	11.71	0.2060	13.88	15.76	19.71
		r = 1	1.00	0.7917	5.47	6.79	9.73
	2 (HQ, SC)	r = 0	7.05	0.6575	13.88	15.76	19.71
		r = 1	2.64	0.4061	5.47	6.79	9.73
S&L with shift dummy[b]	3 (AIC, FPE)	r = 0	10.10	0.0448	8.18	9.84	13.48
	2 (HQ, SC)	r = 0	6.26	0.2135	8.18	9.84	13.48

Note: Deterministic terms in all models: constant, linear trend. Sample period: 2002 M1 - 2011 M12 (including pre sample values). Critical Values from Lütkepohl and Saikkonen (2000).

[a] Shift(s): [2009 M1] [2002 M7].

[b] Trend orthogonal to cointegration relation.

Table A.10

Test for cointegrating rank for (*pwa*, *pccoth*)

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L	7 (AIC)	r = 0	11.71	0.2060	13.88	15.76	19.71
		r = 1	0.15	0.9874	5.47	6.79	9.73
	3 (FPE, HQ, SC)	r = 0	14.03	0.0948	13.88	15.76	19.71
		r = 1	1.04	0.7803	5.47	6.79	9.73
S&L with shift dummy[a]	7 (AIC)	r = 0	12.14	0.1798	13.88	15.76	19.71
		r = 1	1.21	0.7364	5.47	6.79	9.73
	3 (FPE, HQ, SC)	r = 0	13.08	0.1317	13.88	15.76	19.71
		r = 1	0.35	0.9539	5.47	6.79	9.73
S&L with shift dummy[b]	7 (AIC)	r = 0	8.72	0.0803	8.18	9.84	13.48
	3 (FPE, HQ, SC)	r = 0	9.18	0.0662	8.18	9.84	13.48

Note: Deterministic terms in all models: constant, linear trend. Sample period: 2002 M1 - 2011 M12 (including pre sample values). Critical Values from Lütkepohl and Saikkonen (2000).

[a] Shift(s): [2009 M1] [2004 M1].

[b] Trend orthogonal to cointegration relation.

Table A.11

Test for cointegrating rank for (*pcfb*, *m1*, *pwa*)

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L	4 (AIC, FPE)	r = 0	24.26	0.1590	26.07	28.52	33.50
		r = 1	5.38	0.8363	13.88	15.76	19.71
		r = 2	2.03	0.5321	5.47	6.79	9.73
	1 (HQ, SC)	r = 0	22.17	0.2568	26.07	28.52	33.50
		r = 1	2.68	0.9884	13.88	15.76	19.71
		r = 2	0.01	0.9998	5.47	6.79	9.73
S&L with shift dummy[a]	4 (AIC, FPE)	r = 0	32.67	0.0133	26.07	28.52	33.50
		r = 1	2.34	0.9936	13.88	15.76	19.71
		r = 2	0.60	0.8972	5.47	6.79	9.73
	1 (HQ, SC)	r = 0	28.38	0.0521	26.07	28.52	33.50
		r = 1	4.03	0.9387	13.88	15.76	19.71
		r = 2	0.37	0.9493	5.47	6.79	9.73

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Table A.11 (*continued*)

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L with shift dummy[b]	4 (AIC, FPE)	r = 0	19.67	0.0745	18.67	20.96	25.71
		r = 1	2.87	0.6517	8.18	9.84	13.48
	1 (HQ, SC)	r = 0	6.21	0.9400	18.67	20.96	25.71
		r = 1	4.42	0.4093	8.18	9.84	13.48

Note: Deterministic terms in all models: constant, linear trend and seasonal dummies. Sample period: 2002 M1–2011 M12 (including pre sample values). Critical Values from Lütkepohl and Saikkonen (2000).

[a] Shift(s): [2002 M7] [2009 M1].

[b] Trend orthogonal to cointegration relation.

Table A.12
Unit root tests in the presence of structural shift for *w1*.

Shift function	Test Statistic	Lags (1st diff)	Break date	Critical values		
				10%	5%	1%
$f_t^{(1)'}\gamma$	−1.8026	1 (HQ, SC)	2006 M1	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−1.7619	1 (HQ, SC)	2006 M1			
$f_t^{(3)}(\theta)'\gamma$	−1.3243	1 (HQ, SC)	2006 M1			
$f_t^{(1)'}\gamma$	−0.9201	4 (AIC, FPE)[a]	2006 M1			
$f_t^{(2)}(\theta)'\gamma$	−0.8686	4 (AIC, FPE)[a]	2006 M1			
$f_t^{(3)}(\theta)'\gamma$	0.0558	4 (AIC, FPE)	2006 M1			

Note: Time trend included. Critical values (Lanne et al., 2002).

[a] Results when performing the criteria for an unrestricted VAR. For more details on how to run the unit root tests in the presence of structural shift see footnote in Table A.1.

Table A.13
Unit root tests in the presence of structural shift for *w2*

Shift function	Test Statistic	Lags (1st diff)	Break date	Critical values		
				10%	5%	1%
$f_t^{(1)'}\gamma$	−2.9319	1 (HQ, SC)	2006 M1	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−2.7783	1 (HQ, SC)	2006 M1			
$f_t^{(3)}(\theta)'\gamma$	−2.4232	1 (HQ, SC)	2006 M1			

Note: Time trend included. Critical values (Lanne et al., 2002). *Results when performing the criteria for an unrestricted VAR. For more details on how to run the unit root tests in the presence of structural shift see footnote in Table A.1.

Table A.14
Unit root tests in the presence of structural shift for *w3*

Shift function	Test Statistic	Lags (1st diff)	Break date	Critical values		
				10%	5%	1%
$f_t^{(1)'}\gamma$	−2.2393	2 (HQ, SC)	2007 M1	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−2.2298	2 (HQ, SC)	2007 M1			
$f_t^{(3)}(\theta)'\gamma$	−2.3661	2 (HQ, SC)	2007 M1			
$f_t^{(1)'}\gamma$	−1.8745	4 (AIC, FPE)[a]	2007 M1			
$f_t^{(2)}(\theta)'\gamma$	−1.8324	4 (AIC, FPE)[a]	2007 M1			
$f_t^{(3)}(\theta)'\gamma$	−2.3055	4 (AIC, FPE)	2007 M1			

Note: Time trend included. Critical values (Lanne et al., 2002).

[a] Results when performing the criteria for an unrestricted VAR. For more details on how to run the unit root tests in the presence of structural shift see footnote in Table A.1.

Table A.15
Test for cointegrating rank for (*w1*, *w2*, *w3*, *m1*, *pwa*).

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L	2 (AIC, FPE)	r = 0	89.66	0.0001	62.45	66.13	73.42
		r = 1	46.01	0.0423	42.25	45.32	51.45
		r = 2	19.61	0.4205	26.07	28.52	33.50
		r = 3	10.39	0.3041	13.88	15.76	19.71
		r = 4	0.61	0.8947	5.47	6.79	9.73
	1 (HQ, SC)	r = 0	134.98	0.0000	62.45	66.13	73.42
		r = 1	67.12	0.0001	42.25	45.32	51.45
		r = 2	24.01	0.1689	26.07	28.52	33.50
		r = 3	7.77	0.5735	13.88	15.76	19.71
		r = 4	2.40	0.4526	5.47	6.79	9.73

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Table A.15 (*continued*)

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values 90%	95%	99%
S&L with shift dummy[a]	2 (AIC, FPE)	r = 0	76.15	0.0051	62.45	66.13	73.42
		r = 1	42.59	0.0930	42.25	45.32	51.45
		r = 2	27.11	0.0751	26.07	28.52	33.50
		r = 3	8.8	0.4557	13.88	15.76	19.71
		r = 4	1.91	0.5588	5.47	6.79	9.73
	1 (HQ, SC)	r = 0	121.27	0.0000	62.45	66.13	73.42
		r = 1	64.29	0.0002	42.25	45.32	51.45
		r = 2	33.87	0.0088	26.07	28.52	33.50
		r = 3	7.10	0.6510	13.88	15.76	19.71
		r = 4	3.81	0.2328	5.47	6.79	9.73
S&L with shift dummy[b]	2 (AIC, FPE)	r = 0	74.92	0.0002	51.10	54.59	61.53
		r = 1	41.87	0.0092	32.89	35.76	41.58
		r = 2	14.08	0.3247	18.67	20.96	25.71
		r = 3	7.78	0.1176	8.18	9.84	13.48
	1 (HQ, SC)	r = 0	120.33	0.0000	51.10	54.59	61.53
		r = 1	62.22	0.0000	32.89	35.76	41.58
		r = 2	26.15	0.0085	18.67	20.96	25.71
		r = 3	4.73	0.3691	8.18	9.84	13.48

Note: Deterministic terms in all models: constant, linear trend. Sample period: 2002 M1 - 2011 M12 (including pre sample values). Critical Values from Lütkepohl and Saikkonen (2000).

[a] Shift(s): [2006 M1] [2007 M1] [2009 M1] [2002 M7].

[b] Trend orthogonal to cointegration relation.

Table A.16
Test for cointegrating rank for (*w1*, *w2*, *w3*, *m1*, *pwa*)

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values 90%	95%	99%
S&L with shift dummy[a]	1 (HQ, SC)	r = 0	128.26	0.0000	62.45	66.13	73.42
		r = 1	63.57	0.0002	42.25	45.32	51.45
		r = 2	28.57	0.0492	26.07	28.52	33.50
		r = 3	6.27	0.7456	13.88	15.76	19.71
		r = 4	1.07	0.7728	5.47	6.79	9.73

Critical Values from Lütkepohl and Saikkonen (2000).

[a] Shift(s): [2007 M1] [2009 M1].

Table A.17
Unit root tests in the presence of structural shift for *riva*.

Shift function	Test Statistic	Lags (1st diff)	break date	Critical values 10%	5%	1%
$f_t^{(1)'}\gamma$	−3.0143	1 (AIC, FPE, HQ)	2004 M10	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−2.3901	1 (AIC, FPE, HQ)	2004 M10			
$f_t^{(3)}(\theta)'\gamma$	−1.9964	1 (AIC, FPE, HQ)	2004 M10			
$f_t^{(1)'}\gamma$	−4.820	0 (SC)	2004 M10			
$f_t^{(2)}(\theta)'\gamma$	−4.4368	0 (SC)	2004 M10			
$f_t^{(3)}(\theta)'\gamma$	−2.8516	0 (SC)	2004 M10			

Note: Time trend included. Critical values (Lanne et al., 2002). *Results when performing the criteria for an unrestricted VAR. For more details on how to run the unit root tests in the presence of structural shift see footnote in Table A.1.

Table A.18
Test for cointegrating rank for (*riva*, *m1*, *pwa*)

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values 90%	95%	99%
S&L	4 (AIC)	r = 0	27.51	0.0670	26.07	28.52	33.50
		r = 1	12.69	0.1502	13.88	15.76	19.71
		r = 2	1.62	0.6299	5.47	6.79	9.73
	2 (FPE)	r = 0	27.72	0.0632	26.07	28.52	33.50
		r = 1	13.75	0.1046	13.88	15.76	19.71
		r = 2	6.63	0.0543	5.47	6.79	9.73
	1 (HQ, SC)	r = 0	49.67	0.0000	26.07	28.52	33.50
		r = 1	18.16	0.0192	13.88	15.76	19.71
		r = 2	3.84	0.2291	5.47	6.79	9.73

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Table A.18 (*continued*)

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L with shift dummy[a]	4 (AIC)	r = 0	33.35	0.0105	26.07	28.52	33.50
		r = 1	12.04	0.1856	13.88	15.76	19.71
		r = 2	3.50	0.2705	5.47	6.79	9.73
	2 (FPE)	r = 0	0.1588	26.07	28.52	33.50	0.1588
		r = 1	0.0315	13.88	15.76	19.71	0.0315
		r = 2	0.1505	5.47	6.79	9.73	0.1505
	1 (HQ, SC)	r = 0	0.1902	26.07	28.52	33.50	0.1902
		r = 1	0.2494	13.88	15.76	19.71	0.2494
		r = 2	0.2732	5.47	6.79	9.73	0.2732
S&L with shift dummy[b]	4 (AIC)	r = 0	19.93	0.0688	18.67	20.96	25.71
		r = 1	11.75	0.0218	8.18	9.84	13.48
	2 (FPE)	r = 0	15.47	0.2356	18.67	20.96	25.71
		r = 1	7.56	0.1287	8.18	9.84	13.48
	1 (HQ, SC)	r = 0	19.18	0.0861	18.67	20.96	25.71
		r = 1	4.04	0.4629	8.18	9.84	13.48

Note: Deterministic terms in all models: constant, linear trend and seasonal dummies. Sample period: 2002 M1 - 2011 M12 (including pre sample values). Critical Values from Lütkepohl and Saikkonen (2000).

[a] Shift(s): [2002 M7] [2009 M1].

[b] Trend orthogonal to cointegration relation.

Table A.19
Unit root tests in the presence of structural shift for *rtot*

Shift function	Test Statistic	Lags (levels)	Break date	Critical values		
				10%	5%	1%
$f_t^{(1)\prime}\gamma$	−2.7289	3 (AIC, FPE)	2004 M10	−2.76	−3.03	−3.55
$f_t^{(2)}(\theta)'\gamma$	−2.5354	3 (AIC, FPE)	2004 M10			
$f_t^{(3)}(\theta)'\gamma$	−2.5023	3 (AIC, FPE)	2004 M10			
$f_t^{(1)\prime}\gamma$	−3.4265	1 (HQ)[a]	2004 M10			
$f_t^{(2)}(\theta)'\gamma$	−2.7033	1 (HQ)[a]	2004 M10			
$f_t^{(3)}(\theta)'\gamma$	−2.5242	1 (HQ)[a]	2004 M10			
$f_t^{(1)\prime}\gamma$	−4.0395	0 (HQ)	2004 M10			
$f_t^{(2)}(\theta)'\gamma$	−3.6956	0 (HQ)	2004 M10			
$f_t^{(3)}(\theta)'\gamma$	−2.4145	0 (HQ)	2004 M10			

Note: Time trend included. Critical values (Lanne et al., 2002).

[a] Results when performing the criteria for an unrestricted VAR. For more details on how to run the unit root tests in the presence of structural shift see footnote in Table A.1.

Table A.20
Test for cointegrating rank for (*rtot*, *m1*, *pwa*)

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L	5 (AIC)	r = 0	27.49	0.0674	26.07	28.52	33.50
		r = 1	8.66	0.4727	13.88	15.76	19.71
		r = 2	1.60	0.6353	5.47	6.79	9.73
	2 (FPE, HQ)	r = 0	40.05	0.0009	26.07	28.52	33.50
		r = 1	15.18	0.0622	13.88	15.76	19.71
		r = 2	6.51	0.0580	5.47	6.79	9.73
	1 (SC)	r = 0	66.50	0.0000	26.07	28.52	33.50
		r = 1	20.45	0.0073	13.88	15.76	19.71
		r = 2	4.03	0.2088	5.47	6.79	9.73
S&L with shift and without seasonal dummies	5 (AIC)	r = 0	22.58	0.2352	26.07	28.52	33.50
		r = 1	15.03	0.0659	13.88	15.76	19.71
		r = 2	7.86	0.0279	5.47	6.79	9.73
	2 (FPE, HQ)	r = 0	17.64	0.5699	26.07	28.52	33.50
		r = 1	10.46	0.2977	13.88	15.76	19.71
		r = 2	4.57	0.1596	5.47	6.79	9.73
	1 (SC)	r = 0	31.86	0.0174	26.07	28.52	33.50
		r = 1	9.79	0.3581	13.88	15.76	19.71
		r = 2	3.90	0.2224	5.47	6.79	9.73

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Table A.20 (*continued*)

Test	Number of lags (levels)	Null Hypothesis	Test Value	P-value	Critical Values		
					90%	95%	99%
S&L with shift dummy[a]	4 (AIC)	r = 0	29.74	0.0345	26.07	28.52	33.50
		r = 1	19.16	0.0127	13.88	15.76	19.71
		r = 2	2.00	0.5384	5.47	6.79	9.73
	2 (FPE, HQ)	r = 0	19.80	0.4072	26.07	28.52	33.50
		r = 1	13.23	0.1253	13.88	15.76	19.71
		r = 2	4.31	0.1814	5.47	6.79	9.73
	1 (SC)	r = 0	22.34	0.2474	26.07	28.52	33.50
		r = 1	10.83	0.2683	13.88	15.76	19.71
		r = 2	3.58	0.2606	5.47	6.79	9.73
S&L with shift dummy[b]	4 (AIC)	r = 0	21.76	0.0387	18.67	20.96	25.71
		r = 1	14.44	0.0065	8.18	9.84	13.48
	2 (FPE, HQ)	r = 0	13.99	0.3317	18.67	20.96	25.71
		r = 1	7.61	0.1262	8.18	9.84	13.48
	1 (SC)	r = 0	17.27	0.1481	18.67	20.96	25.71
		r = 1	4.31	0.4238	8.18	9.84	13.48

Note: Deterministic terms in all models: constant, linear trend and seasonal dummies. Sample period: 2002 M1 - 2011 M12 (including pre sample values). Critical Values from Lütkepohl and Saikkonen (2000).

[a] Shift(s): [2002 M7] [2009 M1] [2004 M10].

[b] Trend orthogonal to cointegration relation.

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